Multifactor Models are Alive and Well

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ABSTRACT

A large number of studies have investigated the cross-section of average returns on common stocks in the United States and have found little relationship with the estimated beta of the single-factor model. This paper tests the joint roles of an overall market factor, and factors related to firm size (market equity) and style (book equity to market equity) in the cross-section of average stock returns in Australia, as there is little evidence available on the asset pricing theory in markets outside the United States. This paper also tests the claim that the size and style effect is the result of seasonal phenomena. We report that the three-factor model largely explains the variation in stock returns in a meaningful pattern. We also observe that size and style factors do a good job throughout the sample period and reject the claim that these effects are due to seasonal phenomena. Our results document that the explanatory power of the three-factor model is not restricted to a limited set of portfolios. Moreover, our findings do not support the data-snooping hypothesis.

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1. Introduction

ASSET PRICING IN THE AUSTRALIAN SETTING is a largely neglected research area. In contrast, Campbell (2000) reports that, for the last 20 years, theoretical and empirical developments in the United States (US) have taken place within a well-established paradigm. However, to date, few studies have investigated whether the US findings carry over to other markets. This is potentially important because the US results may be sample specific, driven by domestic regulatory, market microstructure and institutional arrangements.

Campbell (2000) observes that one of the outstanding issues to be resolved in asset pricing research is the need to refine our understanding of multifactor asset pricing models in explaining the cross-section of equity returns. In advancing the asset pricing debate, the purpose of this paper is to discuss in more detail the multifactor asset pricing model developed by Fama and French (hereafter FF) (1996), with a view to testing the empirical validity of the model for Australia. Specifically, this paper considers the controversial notion that the anomalies identified by FF (1993, 1996) provide evidence to reject the capital asset pricing model (CAPM), but not a broader rational model in which there are multiple risk factors.

Using a multifactor model developed for the Australian setting, we report that the three factors identified seem to explain the variation in stock returns in a meaningful pattern. We also find that the size (market equity ME) and style[1] (book equity to market equity BE/ME) factors do a good job of explaining cross-sectional stock returns throughout the sample period and reject the claim that the size or style effect is a seasonal phenomenon. The results are consistent with the findings of FF (1996) for the US experience, suggesting that the factors chosen are robust for the international setting.

The evidence presented in this paper challenge Campbell's (2000) observation that anomalies can only be described "*parsimoniously using multifactor models in which the factors are chosen atheoretically to fit the empirical evidence*." This study provides evidence that corroborates the usefulness of a multifactor asset pricing model through the investigation of an out-of-sample check of performance.

The remainder of the paper is organised as follows. In the next section, a brief review of literature is presented. In Section 3 we discuss the applicability of the asset pricing debate for Australia's national system of mandated retirement funding through the employment sector, termed superannuation. Section 4 considers data collection and portfolio construction procedures, with Section 5 outlining the methodology adopted. Empirical evidence is presented in Section 6 and Section 7 concludes the article with a discussion of the implications of the results and avenues for future research.

[1] A value style firm is denoted by a high BE/ME ratio, with growth firms characterized by a low BE/ME ratio.

2. Literature Review

A large number of empirical studies over the past two decades have demonstrated evidence against the prediction of the Sharpe[2] (1964), Lintner (1965) and Black (1972) CAPM. Campbell (2000) observes that, during the 1980s and 1990s, researchers began to look at other characteristics of stocks apart from their betas. This change in research direction led to the identification of puzzles or 'anomalies' that are not captured by systematic risk. The identification of deviations from CAPM has been the catalyst for a re-evaluation and re-assessment of asset pricing technology by researchers leading to the development of multifactor models.

Research by Chan, Hamao and Lakonishok (1991) provided preliminary evidence that firm characteristics (such as size, style, earnings multiples, cash flow and past sales growth ratios) explain the cross-section of average stock returns with greater precision than the single-factor CAPM. FF (1992) examined all of these variables simultaneously for NYSE, AMEX and NASDAQ stocks, making two observations about the cross-section of average stock returns. First, that there is only a weak positive relation between average returns and systematic risk over the period 1941 through 1990 and virtually no relationship over the period 1963 through 1990. Second, that size and style factors capture the cross-section variation in average returns from 1963 through 1990.

In a subsequent study, FF (1996) introduce a multifactor asset pricing model to capture anomalies that are not explained by the CAPM. The research undertakes several tests that suggest a firm's size and style are in fact proxies for the loading on priced risk factors. FF (1996) document covariation in returns related to size and style beyond the covariation explained by the market return, suggesting the existence of a common risk factor. In addition, the loading on zero cost factor portfolios, formed on size and style ratios in concert with a value weighted market portfolio, appear to explain the excess returns of size and style sorted portfolios. The conclusions reached by FF (1996) were consistent with a multifactor version of Merton's (1973) intertemporal capital asset pricing model (ICAPM) or the arbitrage pricing model (APT) of Ross (1976), in suggesting that the higher average returns on value stocks are compensation for risk missed by the CAPM.

The contribution of FF (1996), particularly relating to the interpretation of the results, ignited a flurry of responses from the academy. The first wave of responses, led by DeBondt and Thaler (1987), Haugen (1995) and Lakonishok, Shleifer, and Vishny (1994), suggests that the high returns associated with value stocks are generated by investors who incorrectly extrapolate the past earnings growth rates of firms. As a result, the market tends to undervalue distressed stocks and overvalue growth stocks.
A second area of controversy relates to the controversial findings of Daniel and Titman (1997), forwarding that the return premia on small capitalisation and value stocks does not arise because of the co-movement of these stocks with pervasive factors. Daniel and Titman (1997) argue that

[2] Significant contributions have been made by Banz (1981), Basu (1983), Kandel and Stambaugh (1996), Kothari, Shanken and Sloan (1995), Narasimhan and Titman (1993), Jagannathan and Wang (1996) and Knez and Ready (1997).

it is characteristics rather than the covariance structure of returns that appear to explain the cross-section variation in stock returns. Moreover, Daniel and Titman (1997) claim that the value premia is determined by value characteristics, not risk. Further, Black (1993), MacKinlay (1995) suggest that style phenomenon was obtained for a specific sample and the result is unlikely to be observed out of this sample.

In replying to the critique of the mulitfactor model, Davis, Fama and French (2000) extend data back to 1926 (a 68-year sample period) and expand the sample coverage to all NYSE industrial firms. The results provide no evidence of a sample-specific explanation for the value premium, with the value premium in pre-1963 returns close to that observed for the subsequent period in earlier work. These results led Davis et al (2000) to conclude that the model of FF (1996) explains the value premium better than the characteristic based model of Daniel and Titman (1997). Moreover, Davis et al suggest that the evidence in favour of the characteristic model, provided by Daniel and Titman (1997), appears to be a feature of the sample period.

Further evidence to support the foundations of the received asset pricing theory is provided by Berk (2000). Berk (2000) argues that by sorting data into groups on variables known *a priori* to be correlated with equity returns the explanatory power of a correctly specified asset pricing model can be reduced. Therefore, the explanatory power of the model will always be smaller within a group than in the whole sample. Berk (2000) suggests that a methodology that sorts stocks into more groups, such as the one adopted by Daniel and Titman (1997), destroys the within group explanatory power of a correctly specified asset pricing model.

While the headline debate in asset pricing continues to focus on characteristics, covariances and the sorting of data into groups, a further matter of importance for researchers relates to the turn of the year effect. Although a rich literature exists providing evidence on stock returns being predictable, asset pricing research has paid particular attention to the January effect[3]. In a market characterised by informational efficiency, predictable patterns in stock returns (such as turn of the year effect, day of the week effect or patterns in intraday returns) should not be evident. The difficulty in examining seasonal anomalies is that researchers are ambivalent between whether such anomalies result from a misspecified asset pricing model or informational inefficiency in markets.

Branch (1977) proposes a rule for investors interested in taking advantage of tax selling hypothesis, arguing that investors tend to engage in tax selling towards the end of the year to establish losses on stocks that have declined. The hypothesis suggests that in January, investors reacquire these stocks or other stocks that are attractive. Branch (1977) suggests that this behaviour is responsible for downward pressure in November and December and upward pressure in January.

Brown, Keim, Kleidon, and Marsh (1983) test the January effect for Australia and observe significant abnormal returns in January although the year-end for tax purposes in Australia is

[3] See Dyl (1977), Rozeff and Kinney (1981), Reinganum (1983), Roll (1983), Keim (1983), Kato and Shallheim (1985), Keim and Stambaugh (1984), Brailsford and Easton (1991) and Brailsford (1993) for a complete discussion on January effect.

June 30. Their results indicate large seasonals both in January and June. Although the June results are supported in the sense of tax-loss selling hypothesis, the January effect cannot be explained in Australia.

The sample investigated in this study permits us to deal with the issue of seasonal anomalies. The importance of this area for the real sector is summarised by Fama (1991) stating that "*until we know more about the pricing (and economic fundamentals) of small stocks, inferences should be cautious for the many anomalies where small stocks play a large role.*"

The objective of this paper is to consider two specific research questions. First, we investigate the controversial question of whether value stocks outperform growth stocks. Specifically, we ask whether the model of FF (1996) explains returns for non-US portfolios. Davis et al (2000) note that "*the acid test of a multifactor model is whether it explains differences in returns.*" This paper tests the robustness of the multifactor model on stock returns in a small, open economy that has a vastly different institutional structure than the US.

Second, we consider the efficient market hypothesis in light of the joint-hypothesis problem. The discovery of predictable patterns in security returns, such as seasonal patterns, challenge the validity of an informationally efficient stock market. The paper attempts to provide positive insights into the joint-hypothesis problem through an examination of the seasonal anomaly in the Australian setting where thin evidence is available on this serious issue.

In summary, we suggest that the validity of the FF (1996) model, and the subsequent interpretation of results, can only be settled by testing the three-factor model in different time periods and for different countries. In addition, we also examine the January and July effect (since Australia has a June tax year-end) to determine whether the multifactor model is month specific. These two research questions are addressed to further our understanding of the behaviour of security returns from both a time series and cross-sectional perspective.

3. Institutional Setting

The debate surrounding asset pricing is of particular relevance to Australia, in particular, to superannuation trustees and fund members. The key characteristic that differentiates the evolution of superannuation funds from, say, US pension funds, is that the Australian system is characterised by uniformity, with involuntary participation by employees, employers and government. Moreover, as distinct from other pension funding arrangements internationally, the Commonwealth Government of Australia's regulatory approach to superannuation has been to ensure that the prime responsibility for the viability and prudent operation of the superannuation industry rests with trustees.

The catalyst for the development of Australia's contemporary superannuation regime occurred in 1986 when the Australian Conciliation and Arbitration Commission accepted that claims for superannuation could be part of its conciliation on wages. In entering the wage fixation process, superannuation became the Commonwealth Government of Australia's preferred vehicle for retirement savings.

A second defining event occurred as part of the 1991 Commonwealth budget. During this period, the Government announced its intention to introduce a Superannuation Guarantee Levy (SGL) which required all employers to provide a prescribed level of superannuation support to virtually all of their employees. On 1 July 1992, the SGL commenced at 4 per cent of annual national payroll for companies, scaling up to the 8 per cent of annual national payroll as at 1 July 2000.

The introduction of the award superannuation and the SGL has seen resulted in the greatest institutional change in Australian financial market history. The Australian Prudential Supervision Authority (APRA) (1999) report that, as a result of these two retirement saving initiatives, superannuation assets have increased more then ten-fold from AUD 40 billion on 30 June 1985 to in excess of 409 billion at 30 June 1999. In terms of the broader macroeconomic setting, APRA (1999) report that superannuation assets currently stand at 58 per cent of Australia's gross domestic product.

The institutional setting for retirement funding arrangements in Australia provides a unique sample to test the international findings regarding the rigour of mulitfactor asset pricing models to explain the cross-sectional structure of stock returns. The rapid expansion of superannuation assets over the sample period has resulted in Australia's financial markets, particularly the stock market, being characterised by an unprecedented participation by professional investment managers. It is timely, given the structural change in the Australian market, to undertake an empirical evaluation of multifactor asset pricing models to provide trustees with positive insights into the trade-off between risk and return.

4. Data Selection and Portfolio Construction Procedures

Primark Australia[4], a key independent measurement service in Australia, was commissioned to provide monthly stock returns and accounting data from June 1985 through June 2000. The sample is complete in the sense that it contains no missing data and has been maintained by the same independent data collection agency through the period. As discussed in Section 3 of the paper, the rationale for the commencement date of the sample reflects the introduction of mandated retirement funding through the employment sector in Australia.

The fiscal year end for the majority of Australian firms is June although some firms have a December year-end. We exclude firms with December year-end for portfolio construction purposes. We focus on six portfolios formed yearly from a simple sort of firms into two portfolios of size and three style portfolios. We use all firms, financial and non-financial, for the purpose of this study.

The study by FF (1992) excludes financial firms, as they were interested in studying the explanatory power of leverage on security returns. However, they observed that the leverage effect was not more powerful than the beta. They also observed that the size and distressed firm effect subsumed the explanatory power of all other variables. It is also important to note that the

[4] Primark is a global information services company. We used Primark Australia's 'Datastream' tool to obtain the data for this study.

term leverage does not mean the same for a financial and a non-financial firm. Moreover, there is no reason to expect that size and distressed firm effect have different meanings for a financial and a non-financial firm.

Barber and Lyon (1997) test the size and style returns of financial firms, excluded from the original sample used by FF (1992). Barber and Lyon (1997) find the results from the multifactor model are similar to non-financial firms and suggest that data snooping and selection biases cannot explain the results of FF (1992). Moreover, Barber and Lyon (1997) support the view that the size effect and distressed firms effect do not have different meanings for financial and non-financial firms.

To permit the study to make positive insights into the ability (or otherwise) of multifactor models to explain the cross-section of stock returns in Australia, the research design of FF (1993) is adopted to create size and style portfolios. We define a firm's size or ME as the closing price times the number of shares outstanding at the end of June in year *t*. The style or BE/ME of a firm is the ratio between the book value of stockholders' equity divided by market equity.

In June of each year *t* commencing 1985 to 1999, all stocks in the sample are ranked on size (closing price times number of shares outstanding). The median size is then used to split stocks into two groups of size (small and big). Firms with market equity below the median market value are allocated to the small group and firms with market equity above the median market value are allocated to the big group.

We also independently sort all stocks into three style groups (low, medium, and high) based on the median BE/ME. Firms with BE/ME less than 33.33 percent of median BE/ME are allocated to low BE/ME group, firms with BE/ME greater than 66.67 percent of the median BE/ME are allocated to high BE/ME group and firms with BE/ME between 33.33 and 66.67 percent of the median BE/ME are allocated to medium BE/ME group. We form two groups of size and three style groups. This is done to be consistent with the findings of FF (1993) who suggest that a firm's style has a stronger role in explaining average returns than size. FF (1993) also contend that the splits between groups are arbitrary and results should not be sensitive to these choices.

The BE/ME ratio used to form portfolios in June of year t is then book common equity for the fiscal year ending in calendar year *t-1* divided by the market equity at the end of December of *t-1*. The six portfolios formed at the intersection of size and style are *(S/L, S/M, and S/H; B/L, B/M, and B/H).*

5. Methodology

We investigate the relationship between stock returns, overall market factor, size (ME), and style (BE/ME) by employing the following model:

$$R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \, \mathrm{SMB} + h_i \, \mathrm{HML} + \varepsilon_i \qquad (1)$$

R_{pt} is the average return of a certain portfolio *(S/L, S/M and S/H; B/L, B/M and B/H)***.** R_{ft} is the risk free rate observed at the beginning of the month. Value-weighted monthly portfolio returns are calculated from July to following June. SMB (Small minus Big) is the difference each month between the average of the returns of the three small stock portfolios *(S/L, S/M, and S/H)* and the average of the returns of the three big stock portfolios *(B/L, B/M, and B/H)***.** HML (High minus Low) is the difference between the average of the returns on two high BE/ME *(S/H and B/H)* portfolios and the average of the returns of two low BE/ME *(S/L and B/L)* portfolios. The loadings b_i, s_i, and h_i, are the slopes in the time-series regression.

The CAPM states that the expected returns are a linear function of their market betas, which sufficiently describes the cross-section of expected stock returns. It also states that the market portfolio of invested wealth is mean variance efficient and there should be no added return for bearing nonmarket risk. In short, it is the systematic risk or beta of assets that drives average returns in the CAPM world, with stock returns being unpredictable.

However, recent international evidence suggests that average returns on common equity are related to characteristics such as size, style, earnings to price, cash flow to price, past sales growth, long term past return and short term past return. Since these factors are not captured by the CAPM, they are called anomalies. Cochrane (1999) observes that the identification of anomalies has resulted in researchers deducing that "*we once thought that the CAPM provided good description of why average returns on some stocks, portfolios, funds, or strategies were higher than others. Now we recognise that the average returns of many investment opportunities cannot be explained by the CAPM and 'multifactor models' are used in its place*."

Following the observation of Cochrane (1999), this paper tests the claim that multifactor models dominate the CAPM in that the returns on portfolios are predictable as variables such as size and style explain significant amount of variation in the cross-section of expected stock returns. We also test the existence of a seasonal effect (January and July) by separately analysing the returns of size and book to market equity portfolios. The returns for January, July, non-January and non-July months are separately tested using equation (1).

FF (1992) forward that it is standard in tests of asset pricing models to look for an unexplained January effect. They observe that the January return declines monotonically with increasing size when controlled for style and that the January return increases monotonically with increasing BE/ME when controlled for size. We test for both the January and July effect, as the majority of Australian firms have June as their fiscal year end. If the tax loss-selling hypothesis holds well in Australia we should see a July effect and not a January effect.

6. Empirical Results

A. Tests of the multifactor asset pricing model

Our first research question was to investigate whether a multifactor asset pricing model largely explains the cross-section of average stock returns. Specifically, this study is interested in whether value stocks outperform growth stocks using equation (1). The mean portfolio results and regression results are reported for the sample in Table 1.

Table 1
Panel A: Summary Statistics
Mean Monthly Returns for Australia
Period: 30/06/85 to 30/06/00

	RPTRFT	RMRFT	SMB	HML
S/ L	4.9532 (11.3999) [6]	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)
S/M	1.9800 (9.6497)	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)
S/H	1.7032 (9.5209)	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)
B/L	2.2503 (8.8076)	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)
B/M	1.7231 (8.8724)	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)
B/H	1.0333 (9.6552)	1.2577 (8.5652)	1.2099 (5.1928)	-2.1395 (5.4364)

Table 1-Continued
Panel B: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \text{SMB} + h_i \text{HML} + \varepsilon_{it}$
Period: 30/0685 to 30/06/00

	a	b	s	h	R^2
S/ L	0.984 (2.625) [7]	1.005 (24.554)	0.891 (11.777)	-0.760 (-10.697)	0.84
S/M	0.378 (0.948)	0.940 (21.575)	0.898 (11.159)	0.312 (4.129)	0.74
S/H	0.804 (3.193)	0.995 (36.171)	0.899 (17.695)	0.674 (14.109)	0.89
B/L	0.760 (4.131)	0.988 (49.192)	-0.007 (-2.139	-0.160 (-4.598)	0.93
B/M	0.633 (3.563)	0.983 (50.652)	-0.155 (-4.308)	-0.001 (-0.555)	0.94
B/H	0.774	0.969	-0.007	-0.405	0.84

[6] Standard Deviations in parentheses.

[7] t-statistics in parentheses. We tested the coefficients at 1per cent level of significance.

	(2.468)	**(28.311)**	**(-1.220)**	**(6.814)**	

Our results of the whole sample reveal that the overall market factor (*b* coefficient) is positive and highly significant for all six portfolios. We observe that the size factor (*s* coefficient) is positive and highly significant for the three small portfolios *(S/L, S/M and S/H).* The style factor (*h* coefficient) is negative, but statistically significant for the *(S/L)* portfolio and increases monotonically for the *(S/M)* and *(S/H)* portfolios. The *s* coefficient for the three big portfolios *(B/L, B/M and B/H)* is negative and is significant only for the *(B/M)* portfolio. The *h* coefficient is negative and significant for *(B/L and B/H)* portfolios and not significant for *(B/M)* portfolio.

We answer the question of whether value stocks outperform growth stocks in the affirmative, finding a value premium in Australia. Table 1 provides corroborating evidence of the results from the US by FF (1996) and Davis et al (2000).

In essence, we find that the overall market factor is highly significant in all cases, with the *s* coefficient for the three small portfolios positive in the sense that small firms outperform big firms and the *h* coefficient for the small portfolios increases monotonically. Our results are consistent with the findings of FF (1996) in that Australian stock returns are predictable and the three-factor model seems to explain the variation in the cross-section of average stock returns. Given these results, we next consider question of whether the three-factor model is robust throughout the sample period and is not subject to a seasonal effect.

B. Tests of the seasonal effect

Our second research question considers the seasonal effect (both January and July) to validate the claim of FF (1996) that the multifactor model is robust throughout the sample period. Table 2 reports mean monthly returns and regression results excluding the January month.

Table 2
Panel A: Summary Statistics
Mean Monthly Returns for Australia (Excluding January Month)
Period: 30/06/85 to 30/06/00

	RPTRFT	RMRFT	SMB	HML
S/ L	4.6935 (11.4781)[8]	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)
S/M	1.8564 (9.9322)	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)
S/H	1.3498 (9.6821)	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)
B/L	2.2785 (9.0517)	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)
B/M	1.7129 (9.1331)	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)
B/H	0.7122 (9.8229)	1.1759 (8.8152)	1.0654 (5.2091)	-2.3499 (5.4937)

[8] Standard Deviations in parentheses.

Table 2-Continued
Panel B: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \text{ SMB} + h_i \text{ HML} + \varepsilon_{it}$
Period: 30/06/85 to 30/06/00

	a	b	s	h	R^2
S/ L	**0.749 (1.913)**[9]	**0.995 (24.084)**	**0.837 (10.664)**	**-0.802 (-10.977)**	**0.84**
S/M	**0.550 (1.277)**	**0.946 (20.826)**	**0.924 (10.703)**	**0.336 (4.185)**	**0.74**
S/H	**0.841 (3.119)**	**0.991 (34.850)**	**0.873 (16.152)**	**0.675 (13.421)**	**0.89**
B/L	**0.869 (4.492)**	**0.988 (48.397)**	**-0.009 (-2.395)**	**-0.148 (-4.100)**	**0.93**
B/M	**0.689 (3.625)**	**0.984 (49.078)**	**-0.155 (-4.064)**	**-0.001 (-0.384)**	**0.94**
B/H	**0.582 (1.783)**	**0.961 (27.885)**	**-0.120 (-1.828)**	**0.371 (6.092)**	**0.85**

We find that the results of the sample excluding the January month are similar to the whole sample in that the overall market factor is highly significant for all six portfolios and the s coefficient is positive and significant for the three small portfolios *(S/L, S/M and S/H)*. The *s* coefficient for the three big portfolios *(B/L, B/M and B/H)* is negative and is significant only for *(B/M)* portfolio. The behaviour of the *h* coefficient in the ex-January sample is similar to the h coefficient of the whole sample. The *h* coefficient is negative for *(B/L and B/M)* portfolios and positive for *(B/H)* portfolio. However, the *h* coefficient for the *(B/M)* portfolio is not statistically significant.

Table 3
Panel A: Summary Statistics
Mean Monthly Returns for Australia (Only January Month)
Period: 30/06/85 to 30/06/00

[9] t-statistics in parentheses. We tested the coefficients at 1per cent level of significance.

	RPTRFT	RMRFT	SMB	HML
S/ L	**7.8101 (10.4286)**[10]	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**
S/M	**3.3396 (5.7266)**	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**
S/H	**5.5908 (6.5728)**	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**
B/L	**1.9399 (5.6510)**	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**
B/M	**1.8354 (5.4158)**	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**
B/H	**4.5652 (6.8584)**	**2.1578 (5.1730)**	**2.80000 (4.8934)**	**0.1747 (4.2503)**

Table 3-Continued
Panel B: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \text{ SMB} + h_i \text{ HML} + \varepsilon_{it}$

Period: 30/06/85 to 30/06/00

	a	b	s	h	R^2
S/ L	**0.988 (0.573)**[11]	**1.244 (5.207)**	**1.491 (4.194)**	**0.209 (-0.513)**	**0.86**
S/M	**-0.678 (-0.545)**	**0.840 (4.876)**	**0.777 (3.030)**	**0.175 (0.596)**	**0.75**
S/H	**-0.556 (-0.703)**	**1.071 (9.774)**	**1.321 (8.107)**	**0.778 (4.159)**	**0.92**
B/L	**-0.895 (0.713)**	**1.013 (10.255)**	**0.241 (1.639)**	**-0.146 (-0.863)**	**0.91**
B/M	**-0.212 (-0.281)**	**0.991 (9.484)**	**-0.003 (-0.229)**	**0.004 (0.273)**	**0.90**
B/H	**0.861 (0.559)**	**1.151 (5.392)**	**0.383 (1.208)**	**0.841 (2.307)**	**0.74**

The results of the January month alone, provided in Table 3, indicate that the overall market factor is significant for all six portfolios and the *s* coefficient is positive and significant for *(S/L and S/H)* portfolios. The *h* coefficient for the *(S/L and S/M)* portfolios are not significant, but significant for the *(S/H)* portfolio. The *s* and *h* coefficients for the three big portfolios *(B/L, B/M and B/H)* are not statistically significant.

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[10] Standard Deviations in parentheses.

[11] t-statistics in parentheses. We tested the coefficients at 1per cent level of significance.

We do observe that the mean monthly returns are positive and significantly higher for the three small portfolios *(S/L, S/M and S/H),* (Table 3) when compared with the whole sample (Table 1). We also find that the mean monthly returns for the *(B/H)* portfolio in the January month sample is higher than the whole sample. While these observations provide some evidence of anomalous behaviour, we conclude that as the *s* coefficient is not significant for the majority of portfolios *(S/M, B/L, B/M and B/H)* and *h* for none of the six portfolios, we reject any claim that the size or style effect is a January phenomenon in the Australian setting.

Table 4
Panel A: Summary Statistics
Mean Monthly Returns for Australia (Excluding July Month)
Period: 30/06/85 to 30/06/00

	RPTRFT	RMRFT	SMB	HML
S/ L	4.0878 (11.1807)[12]	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)
S/M	1.4446 (9.5524)	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)
S/H	0.9893 (9.3171)	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)
B/L	1.6925 (8.7574)	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)
B/M	1.2293 (8.8721)	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)
B/H	0.7197 (9.8499)	0.8138 (8.5979)	0.9601 (5.1310)	-1.9491 (5.5463)

[12] Standard Deviations in parentheses.

Table 4-Continued
Panel B: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \text{ SMB} + h_i \text{ HML} + \varepsilon_{it}$
Period: 30/06/85 to 30/06/00

	a	b	s	h	R^2
S/ L	**0.933 (2.445)**[13]	**1.013 (23.618)**	**0.912 (11.381)**	**-0.746 (-10.337)**	**0.83**
S/M	**0.422 (1.043)**	**0.950 (20.848)**	**0.893 (10.497)**	**0.312 (4.073)**	**0.73**
S/H	**0.673 (2.928)**	**0.980 (37.890)**	**0.930 (19.265)**	**0.705 (16.204)**	**0.91**
B/L	**0.674 (3.745)**	**0.986 (48.711)**	**-0.005 (-1.460)**	**-0.138 (-4.058)**	**0.94**
B/M	**0.574 (3.250)**	**0.977 (49.219)**	**-0.148 (-3.993)**	**-0.001 (-0.040)**	**0.94**
B/H	**0.778 (2.294)**	**0.966 (26.549)**	**-0.120 (-1.367)**	**0.394 (6.343)**	**0.85**

We now turn our attention to the July anomaly. The results of the sample excluding July month (Table 4) indicates that the overall market factor is highly significant for all six portfolios and the s coefficient for the three small portfolios *(S/L, S/M and S/H)* are positive and highly significant. The *h* coefficient for the *(S/L)* portfolio is negative but significant and positive and significant for both *(S/M and S/H)* portfolios. The *s* coefficient for the three big portfolios *(B/L, B/M and B/H)* is negative and is significant only for the *(B/M)* portfolio. The *h* coefficient for *(B/L and B/M)* is negative and for *(B/H)* positive. We also observe that the *h* coefficient is significant for the *(B/L and B/H)* portfolios. It is important to note that the three coefficients *(b, s and h)* behave similar to the coefficients in the whole sample.

[13] t-statistics in parentheses. We tested the coefficients at 1per cent level of significance.

<u>Table 5</u>
<u>Panel A: Summary Statistics</u>
<u>Mean Monthly Returns for Australia (Only July Month)</u>
Period: 30/06/85 to 30/06/00

	RPTRFT	RMRFT	SMB	HML
S/ L	14.4732 (9.5549)[14]	6.1404 (6.6459)	3.9583 (5.2475)	-4.2340 (3.5215)
S/M	7.8700 (7.8052)	6.1404 (6.6459)	3.9583 (5.2475)	-4.2340 (3.5215)
S/H	9.5560 (8.3683)	6.1404 (6.6459)	3.9583 (5.2475)	-4.2340 (3.5215)
B/L	8.3863 (7.0310)	6.1404 (6.6459)	3.9583 (5.2475)	-4.2340 (3.5215)
B/M	7.1554 (7.0649)	6.1404 (6.6459)	3.9583 (3.9583)	-4.2340 (3.5215)
B/H	4.4825 (6.4414)	6.1404 (6.6459)	3.9583 (3.9583)	-4.2340 (3.5215)

<u>Table 5-Continued</u>
<u>Panel B</u>: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i$ SMB + h_i HML $+ \varepsilon_{it}$
Period: 30/06/85 to 30/06/00

	a	b	s	h	R^2
S/ L	2.641 (1.092)[15]	0.795 (3.733)	0.637 (2.212)	-1.047 (-2.402)	0.76
S/M	-0.448 (-0.165)	0.771 (3.226)	1.016 (3.142)	0.105 (0.214)	0.66
S/H	2.891 (1.252)	1.061 (5.225)	0.323 (1.176)	0.267 (0.643)	0.72
B/L	2.534 (2.236)	0.901 (9.041)	-0.471 (-3.490)	-0.515 (-2.525)	0.90
B/M	1.337 (1.102)	0.942 (8.823)	-0.364 (-2.522)	-0.348 (-1.594)	0.89
B/H	1.214 (0.621)	0.784 (4.557)	-0.188 (-0.810)	0.189 (0.536)	0.66

Mean monthly returns and regression results for the month of July only are presented in Table 5. The results of the July month sample indicate that the overall market factor is positive and

[14] Standard Deviations in parentheses.

[15] t-statistics in parentheses. We tested the coefficients at 1 per cent level of significance.

significant and the s coefficient is positive but not significant for the three small portfolios *(S/L, S/M and S/H)***.** The *s* coefficient is also not significant for *(B/M and B/H)* portfolios. The *h* coefficient is not significant for any of the six portfolios.

We do observe that the mean monthly returns for all the six portfolios *(S/L, S/M, S/H and B/L, B/M and B/H)* are substantially higher for the July month sample when compared specifically with the returns of the June month only sample. The mean monthly returns are negative for all the six portfolios in the June month only sample. In our view, the findings provide limited support to the tax-loss selling hypothesis, as the returns in the June month are negative but become positive and higher in July. However, at the 1 per cent level, we conclude that as the *s* coefficient is not significant for *(S/L, S/M, S/H; B/M and B/H)* and *h* coefficient not significant for any of the six portfolios we reject the claim that size and style factors are a July phenomenon.

Table 6
Panel A: Summary Statistics
Mean Monthly Returns for Australia (Excluding January and July)
Period: 30/06/85 to 30/06/00

	RPTRFT	**RMRFT**	**SMB**	**HML**
S/ L	**3.7156** **(11.2178)**[16]	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**
S/M	**1.2551** **(9.8466)**	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**
S/H	**0.529** **(9.4415)**	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**
B/L	**1.6677** **(9.0225)**	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**
B/M	**1.1687** **(9.1565)**	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**
B/H	**0.3351** **(10.0365)**	**0.6794** **(8.8685)**	**0.7761** **(5.1335)**	**-2.1615** **(5.6268)**

[16] Standard Deviations in parentheses.

<u>Table 6-Continued</u>
<u>Panel B</u>: $R_{pt} - R_{ft} = a_i + b_i (R_{mt} - R_{ft}) + s_i \text{ SMB} + h_i \text{ HML} + \varepsilon_{it}$

Period: 30/06/85 to 30/06/00

	a	b	s	h	R^2
S/ L	**0.673 (1.685)**[17]	**0.999 (22.990)**	**0.847 (10.109)**	**-0.789 (-10.635)**	**0.84**
S/M	**0.616 (1.402)**	**0.959 (20.061)**	**0.925 (10.022)**	**0.338 (4.137)**	**0.74**
S/H	**0.697 (2.838)**	**0.974 (36.445)**	**0.899 (17.435)**	**0.707 (15.477)**	**0.91**
B/L	**0.782 (4.133)**	**0.984 (47.830)**	**-0.006 (-1.749)**	**-0.125 (-3.564)**	**0.94**
B/M	**0.628 (3.845)**	**0.978 (47.454)**	**-0.149 (-3.739)**	**0.003 (-0.109)**	**0.94**
B/H	**0.576 (1.719)**	**0.970 (26.601)**	**-0.110 (-1.570)**	**0.377 (6.047)**	**0.85**

When both January and July are excluded from the sample we observe that the overall market factor is positive and highly significant for all six portfolios *(S/L, S/M/ S/H and B/L, B/M and B/H)*, and the *s* coefficient for the three small portfolios *(S/L, S/M and S/H)* is positive and highly significant. The *h* coefficient for the *(S/L)* portfolio is negative but significant, but positive and significant for *(S/M and S/H)* portfolios. The *s* coefficient for the three big portfolios *(B/L, B/M and B/H)* is negative and is significant only for the *(B/M)* portfolio. However, the *h* coefficient for *(B/L and B/H)* is significant and for *(B/M)* not significant. It is important to note that the coefficients *(b, s and h)* behave very similar to the coefficients of the whole sample.

For completeness, we compare the parameter estimates and t-statistics of the three regressions in Tables 1, 2 and 3 (whole sample, the sample excluding January month and the January month alone) to establish any evidence of a January effect. We do not observe a turn of the year effect, as there is no significant difference between the whole sample results and the results excluding the January month. We also observe that the s coefficient is not significant for the *(S/M)* portfolio and for the three big portfolios in the January month only sample. We also observe that the *h* coefficient is not significant for the *(S/L and S/M)* portfolios and for the three big portfolios. However, it is only the overall market factor *(b coefficient)* that behaves similarly to the coefficient of the whole sample and sample excluding January month results. We therefore, reject the claim that the size and style effects are a January phenomenon. In our view, the model is robust and the three factors do a good job in explaining the variation in average Australian stock returns throughout the sample period regardless of whether January is in the sample or not.

[17] t-statistics in parentheses. We tested the coefficients at 1per cent level of significance.

Given that the majority of Australian firms have a financial year end in June, we also compare the parameter estimates of the three regressions in Tables 1, 4 and 5 (whole sample, the sample excluding the July month and the July month alone) to detect traces of a July effect. We do not observe a July effect, as there is no significant difference between the whole sample results and the results excluding the July month. Importantly, we observe that the *s* coefficient in the July month only sample is not significant for *(S/L, S/M and S/H; B/M and B/H)* portfolios. The *h* coefficient is not significant for any of the six portfolios. We therefore, reject the presence of a July effect since both the coefficients *(s and h)* are not significant for most of the portfolios. We only find the *b* coefficient to be significant for all the six portfolios.

We do not find any difference between the whole sample results and the sample excluding the July month. We also find that *(s and h)* coefficients are not significant for most of the portfolios in the July month only sample. Hence, we reject the existence of a July effect and find that the three-factor model explains the variation in stock returns throughout the sample period.

C. Diagnostics

As the motivation for this study is to provide evidence on the robustness of mulitfactor models through the investigation of an out-of-sample check of performance, the following section reports a synopsis of the diagnostics from running equation (1). We report the t-statistics for statistical significance and restrict the level of significance to 1 percent. Our objective has been to keep the regression model parsimonious. We observe that the market factor, size effect and style effect do a good job in explaining the variation in stock returns in a meaningful pattern.
We are also cognisant of the fact that important variables in terms of relevance should not be excluded just to follow the path of principle of parsimony. We report very high coefficient of determination (R^2) values, which suggest that the explanatory variables explain a large proportion of the variation in the dependent variable. The (R^2) values for our six portfolios ranges between 0.66 to 0.94 which suggests that our independent variables *(RMRFT, SMB and HML)* explain between 66 to 94% of the variation in our dependent variable *(RPTRFT)* across the various cuts at the data.

Table 7
Pearson Correlation Coefficient Matrix for the full sample

Portfolio	RMRFT	SMB	HML
Small to Low			
RMRFT	**1.000 (0.0)**	**-0.196 (0.004)**	**0.074 (0.162)**
SMB	**-0.196 (0.004)**	**1.000 (0.0)**	**-0.455 (0.000)**
HML	**0.074 (0.162)**	**-0.455 (0.000)**	**1.000 (0.0)**
Small to Medium			
RMRFT	**1.000 (0.0)**	**-0.196 (0.004)**	**0.074 (0.162)**
SMB	**-0.196 (0.004)**	**1.000 (0.0)**	**-0.455 (0.000)**
HML	**0.074 (0.162)**	**-0.455 (0.000)**	**1.000 (0.0)**
Small to High			
RMRFT	**1.000 (0.0)**	**-0.196 (0.004)**	**0.074 (0.162)**
SMB	**-0.196 (0.004)**	**1.000 (0.0)**	**-0.455 (0.000)**
HML	**0.074 (0.162)**	**-0.455 (0.000)**	**1.000 (0.0)**

Table 7- Continued
Pearson Correlation Coefficient Matrix for full sample

PORTFOLIO	RMRFT	SMB	HML
Big to Low			
RMRFT	**1.000** **(0.0)**	**-0.196** **(0.004)**	**0.074** **(0.162)**
SMB	**-0.196** **(0.004)**	**1.000** **(0.0)**	**-0.455** **(0.000)**
HML	**0.074** **(0.162)**	**-0.455** **(0.000)**	**1.000** **(0.0)**
Big to Medium			
RMRFT	**1.000** **(0.0)**	**-0.196** **(0.004)**	**0.074** **(0.162)**
SMB	**-0.196** **(0.004)**	**1.000** **(0.0)**	**-0.455** **(0.000)**
HML	**0.074** **(0.162)**	**-0.455** **(0.000)**	**1.000** **(0.0)**
Big to High			
RMRFT	**1.000** **(0.0)**	**-0.196** **(0.004)**	**0.074** **(0.162)**
SMB	**-0.196** **(0.004)**	**1.000** **(0.0)**	**-0.455** **(0.000)**
HML	**0.074** **(0.162)**	**-0.455** **(0.000)**	**1.000** **(0.0)**

We also conduct a correlation analysis of the independent variables, the results of which are reported in Table 7. Additionally we also obtain a plot of the observations, which are not reported for reasons of space, to observe the impact (if any) of influential observations. We do observe a small number of influential observations and hence we test the classical linear regression model (CLRM) assumption of no multicollinearity. We report the collinearity diagnostics in Table 8. We find no trace of multicollinearity and conclude that the correlation does not distract from the regressions of the model.

Table 8
Collinearity Diagnostics for full sample

Portfolio (Small to Low)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

Portfolio (Small to Medium)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

Portfolio (Small to High)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

Finally, we conduct tests of homogeneity and normality to test the CLRM assumption that disturbances (μ_i) appearing in the regression function are homoscedastic. In essence this is the assumption of equal variance. Although the problem of heteroscedasticity is more common in cross-sectional data than in time-series data, we still run tests for heteroscedasticity because of the presence of some influential observations. We do not report the statistics for White's (1980) test for reasons of space, but do not find any evidence of hetroscedasticity in any of the six Size-BE/ME portfolios.

7. Conclusion

This paper tests the explanatory power of the multifactor model in an Australian setting. The evidence suggests that value stocks (high BE/ME) materially outperform growth stocks (low BE/ME). We observe that the overall market factor, size and style coefficients explain the variation in the cross-section of expected stock returns in a meaningful manner. We also note that the three-factor is robust in the sense that the explanatory power of the coefficients is uniform throughout the sample period and is not subject to a seasonal effect.

Our results are consistent with the findings of FF (1996) for US portfolios in providing international evidence that a three-factor model explains the variation in the cross-section of average stock returns. The evidence for the Australian market documents that multifactor models are alive and well. Consequently, we reject the data-snooping hypothesis, arguing that if FF (1996) were involved in data-snooping, our results would have contradicted their findings. In summary, this study provides support for a broader, rational asset pricing model in which there are multiple risk factors.

Our results have immediate implications for Australia's burgeoning investment management industry, specifically in relation to the asset selection problem. Australia's regulatory approach to retirement savings, in both a statutory and common law sense, places responsibility for the prudent management of superannuation assets with the trustee. These responsibilities for the trustees of superannuation funds include the selection and engagement of investment managers. The findings presented in this study raises some important policy questions concerning the asset selection mandates trustees are striking with investment managers.

A specific direction for future research is to test the multifactor model under a wide variety of conditions to determine whether the risk and return characteristics are different for, say, a developed market from a developing market, to ascertain whether the value premium is pervasive. This is an issue we will explore in our next paper. The evidence provided in this paper suggests that the value premium is real and that the trustees of superannuation funds, on behalf of their members, should have a value-bias when considering the asset selection problem.

Table 1

Summary Statistics and Three-Factor Regressions for Simple Monthly Percent Excess Returns on 6 Portfolios Formed on Size and Book to Market Equity Ratio (BE/ME): 7/85 to 6/00, 180 Months

Rft is the Commonwealth Government of Australia 13-week treasury notes rate obtained from the Reserve Bank of Australia. At the end of June of each year t stocks are assigned to two portfolios of size (Small and Big) based on whether their June Market Equity (ME) defined as closing price times Number of shares outstanding is above or below the median ME. The same stocks are allocated in an independent sort to three-book equity to market equity portfolios (Low, Medium, and High) based on the breakpoints for the bottom 33.33 percent and top 66.67 percent. Low portfolios consist of firms with breakpoints less than 33.33 percent of median book to market equity. High portfolios consist of firms with breakpoints more than 66.67 percent of median book to market equity and the balance firms are assigned the medium portfolio.

Six ME-BE/ME portfolios are formed at the intersection of the two size portfolios and three book to market equity portfolios. The six portfolios formed are *(S/L, S/M, and S/H; B/L, B/M, and B/H).* Value-weight monthly returns on the six portfolios are calculated from the following July to June. The explanatory variables **RM, SMB, and HML** are defined as follows: RM (market return) is the value-weight market return on all stocks in the six portfolios and includes the negative Book Equity stocks which were excluded from the sample while forming BE/ME portfolios. SMB (Small minus Big) is the difference each month between the average of the returns of the three small stock portfolios *(S/L, S/M, and S/H)* and the average of the returns of the three big portfolios *(B/L, B/M, and B/H).* HML (High minus Low) is the difference between the average of the returns of the two high BE/ME portfolios *(S/H, B/H)* and the average of the returns on the two low BE/ME portfolios *(S/L, B/L).*

We define Book Equity (BE) as the book value of common shareholder’s equity plus the balance sheet deferred taxes (if any) and minus the book value of preferred stocks. The BE/ME ratio used to form portfolios in December of each year t is the book common equity for the fiscal year ending in calendar year t-1 / market equity at the end of December of t-1. While forming portfolios we exclude negative Book equity firms, as they do not have meaningful explanations. We include firms only with ordinary equity for portfolio construction purposes.

Portfolio (Big to Low)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

Portfolio (Big to Medium)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

Portfolio (Big to High)

Number	Eigen Value	Condition Index	Var Prop Intercept	Var Prop RMRFT	Var Prop SMB	Var Prop HML
1	1.739	1.000	0.12	0.00	0.14	0.16
2	1.128	1.242	0.14	0.61	0.06	0.00
3	0.662	1.621	0.67	0.37	0.18	0.05
4	0.471	1.921	0.08	0.02	0.62	0.79

References

Banz, Rolf W., 1981, The relation between return and market value of common stocks, *Journal of Financial Economics* 9, 3-18.

Basu, Sanjoy, 1983, The relationship between earnings yield, market value, and return for NYSE common stocks: Further evidence, *Journal of Financial Economics* 12, 129-156.

Berk, Jonathan B., 2000, Sorting Out Sorts, *Journal of Finance* 55, 407-427.

Black, Fischer, 1972, Capital market equilibrium with restricted borrowing, *Journal of Business* 45, 444-454.

Black, Fischer, 1993, Beta and return, *Journal of Portfolio Management* 20, 8-18.

Brailsford, Tim J., 1993, Seasonality and institutional factors in the New Zealand Equity Market, *Pacific Accounting Review* 5, 1-26.

Brailsford, Tim J., and S.A. Easton, 1991, Seasonality in Australian Share Price Indices between 1936 and 1957, *Accounting and Finance* 31, 69-85.

Branch, Ben, 1977, A tax loss trading rule, *Journal of Business* 50, 198-207.

Brown, Philip, Donald B. Keim, Allan W. Kleidon, and Terry A. Marsh, 1983, Stock return seasonalities and the tax-loss selling hypothesis, *Journal of Financial Economics* 12, 105-127.

Campbell, John Y., 2000, Asset pricing at the millennium, *Journal of Finance* 55, 1515-1567.

Chan, Louis K.C., Yasushi Hamao, and Josef Lakonishok, 1991, Fundamental and stock returns in Japan, *Journal of Finance* 46, 1739-1789.

Cochrane, John H., 1999, New facts in finance, *Economic Perspectives* 23, 1-31.

DeBondt, Werner F.M., and Richard H. Thaler, 1987, Further evidence on investor overreaction and stock market seasonality, *Journal of Finance* 42, 557-581.

Dyl, Edward A., 1977, Capital gains taxation and year-end stock market behaviour, *Journal of Finance* 32, 165-175.

Fama, Eugene F., and Kenneth R. French, 1992, The cross-section of expected stock returns, *Journal of Finance* 42, 427-465.

Fama, Eugene F., and Kenneth R. French, 1993, Common risk factors in the returns on stocks and bonds, *Journal of Financial Economics* 33, 3-56.

Fama, Eugene F., and Kenneth R. French, 1996, Multifactor explanations of asset pricing anomalies, *Journal of Finance* 51, 55-84.

Fama, Eugene F., and Kenneth R. French, 1998, Value versus growth: The international evidence, *Journal of Finance* 53, 1975-1999.

Haugen, Robert A., 1995, *The New Finance: Evidence against Efficient Markets,* (Prentice Hall, New Jersey).

Jagannathan, Ravi, and Zhenyu Wang, 1996, The Conditional CAPM and the cross-section of expected returns, *Journal of Finance* 51, 3-53.

Kandel, Shmuel, and Robert F. Stambaugh, 1996, On the predictability of stock returns: An asset allocation perspective, *Journal of Finance* 51, 385-424.

Kato, K., and J. Shallheim, 1985, Seasonal and Size anomalies in the Japanese stock market, *Journal of Financial and Quantitative Analysis* 20, 243-260.

Keim, Donald B., 1983, Size related anomalies and stock return seasonality, *Journal of Financial Economics* 12, 13-32.

Keim, Donald B., and Robert F. Stambaugh, 1984, A further investigation of the weekend effect in stock returns, *Journal of Finance* 39, 819-840.

Knez, Peter J., and Mark J. Ready, 1997, On the robustness of size and book-to-market in cross-sectional regressions, *Journal of Finance* 52, 1355-1382.

Kothari, S. P., Jay Shanken, and Richard G. Sloan, 1995, Another look at the cross-section of expected stock returns, *Journal of Finance* 50, 185-224.

Lakonishok, Josef, Andrei Shleifer, and Robert W. Vishny, 1994, Contrarian investment, extrapolation and risk, *Journal of Finance* 49, 1541-1578.

Linter, John, 1965, The valuation of risky assets and the selection of risky investments in stock portfolios and capital budgets, *Review of Economics and Statistics* 47, 13-37.

MacKinlay, A. Craig, 1995, Multifactor models do not explain deviations from the CAPM, *Journal of Financial Economics* 38, 3-28.

Merton, Robert C., 1973, An intertemporal asset pricing model, *Econometrica* 41, 867-887.

Narasimhan, Jegadeesh, and Sheridan Titman, 1993, Returns to buying winners and selling losers: Implications for stock market efficiency, *Journal of Finance* 38, 65-91.

Reinganum, Marc R., 1983, The anomalous stock market behaviour of small firms in January: Empirical tests for tax-loss selling effects, *Journal of Financial Economics* 12, 89-104.

Roll, Richard, 1983, Vas Ist Das?, *Journal of Portfolio Management* 9, 18-28.

Ross, Stephen A., 1976, The arbitrage theory of capital asset pricing, *Journal of Economic Theory* 13, 341-360.

Rozeff, Michael S., and William R. Kinney, 1976, Capital market seasonality: The case of stock returns, *Journal of Financial Economics* 3, 379-402.

Sharpe, William F., 1964, Capital asset prices: A theory of market equilibrium under conditions of risk, *Journal of Finance* 19, 425-442.

White, Halbert, 1980, A heteroscedasticity consistent covariance matrix estimator and a direct test of heteroscedasticity, *Econometrica* 48, 817-838.